Psychometric properties of a German-language scale to assess sport satisfaction

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ABSTRACT

This study presents and psychometrically evaluates a German-language 5-item scale designed to assess satisfaction with practising a sport: the Sport Satisfaction Scale (SSS; Sportzufriedenheitsskala, SZS). It is an adaptation of an existing instrument measuring domain-specific self-satisfaction to the sport context. We analysed the dimensionality of the scale, assessed its internal consistency, tested measurement invariance across a range of demographic and sport-related variables, and explored associations with related psychological constructs and external criteria. The sample comprised 285 young adults (M_{age} = 23.06 years, *SD* = 3.49; 43.9% female, 55.8% male, 0.4% other). The analyses supported the hypothesised one-factor structure of the SSS and revealed satisfactory internal consistency (ω = .80). Measurement invariance analyses provided evidence for weak, (partial) strong, and (partial) strict invariance across gender, participation in organised sport, and participation in competitive sport. There was no indication of monotonic differential item functioning with respect to age and weekly training volume. Finally, the findings aligned with theoretical expectations, showing that sport satisfaction correlated positively with several dimensions of passion, physical self-concept, and athletic identity. In conclusion, the SSS demonstrates sound psychometric properties, supporting its use as an economic and valid research tool in sport and exercise settings.

Keywords

athletes, exercise, satisfaction, measurement, validation

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Introduction

In general psychology, and particularly within positive psychology, satisfaction is considered a key construct, commonly used as an indicator of subjective well-being (Pavot & Diener, 2008). It can be assessed in relation to different aspects, ranging from global life satisfaction to specific life domains such as work or job (e.g., Judge et al., 2001), residential environment (e.g., Emami & Sadeghlou, 2021), romantic relationships (e.g., Bühler & Orth, 2025), family (e.g., Zabriskie & Ward, 2013), financial situation (e.g., Ngamaba et al., 2020), or leisure (e.g., Beard & Ragheb, 1980).

The field of sports and exercise has also been concerned with satisfaction. Many coaches, sports instructors, and physical education teachers intuitively associate satisfaction with continued sport participation, performance, and success, especially given the typically voluntary nature of engagement in sport activities. Accordingly, sport and exercise psychology has studied satisfaction in diverse contexts. Among these, athletes' life satisfaction (and more generally: their well-being) has probably been the focus of research in recent years (e.g., Filbay et al., 2019). Looking further back, athlete satisfaction has been a core construct in social-psychological approaches to leadership in sports and physical education. In particular, the multidimensional model of leadership (Chelladurai, 1993) linked coaching behaviour to team member satisfaction (along with group performance). Notably, the development of an instrument to measure multidimensional athlete satisfaction (Riemer & Chelladurai, 1998) provided an impetus for research. For example, it was also used in a related field to assess the quality of the dyadic relationships that coaches form with their athletes (e.g., Jowett et al., 2012). Similarly, in the tradition of achievement goal theory (Nicholls, 1984) studies on coach-, teacher-, and peer-created motivational climate have dealt with the effects which the sport or physical education setting has on satisfaction with participation in sport (e.g., García-Calvo et al., 2014). Satisfaction in sports and exercise settings is also part of the motivational model of the coach-athlete relationship (Mageau & Vallerand, 2003). Drawing on the basic psychological needs theory within self-determination theory (e.g., Deci & Ryan, 2000), Mageau and Vallerand (2003) describe the psychological processes through which coaching behaviours influence athletes' motivation, thereby also delineating the extent to which individuals partake in sport activities due to the associated pleasure and satisfaction. Finally, although coming from a different tradition, exercise psychology has provided much empirical evidence for the effects of acute exercise and sport on the affective component of well-being (e.g., Reed & Ones, 2006).

Taken together, the existing literature underscores the relevance of satisfaction for psychological functioning and well-being in sport. However, it also reveals that no instrument is currently available for measuring sport satisfaction. By this, we refer to the cognitive – as opposed to the affective – component of subjective well-being (Diener et al., 1999) experienced by individuals who engage in a particular sport with a certain frequency and regularity, potentially even in a competitive context. As a result of this commitment, sport constitutes a significant part in individuals' lives in affective, behavioural, cognitive, and temporal terms, like other life domains such as work, relationships, leisure, and health. Among the instruments used in the literature (e.g., García-Calvo et al., 2014; Gaudreau & Braaten, 2016; Luth et al., 2017; Riemer & Chelladurai, 1998), the adaptation used by, for example, Baudin et al. (2011) or Somoğlu and Cihan (2022), which replaces "life" with "sport life" in the extensively studied Satisfaction With Life Scale (SWLS; Diener et al., 1985), appears to be promising for two reasons: (1) the abstract wording enables application across populations practising different sports, and (2) with only five items, the scale can be readily incorporated into longer surveys. Test economy is relevant in situations where sport satisfaction is not the focus of a study but rather a peripheral variable. However, although adapting scale items from one context to another is common in research, it is not without its problems (e.g., Heggestad et al., 2019). This is also likely to be the case when the contexts differ in terms of specificity,

in particular, when items measuring global satisfaction are adapted to assess domain-specific satisfaction. Indeed, applying the proposed change to the SWLS item "The conditions of my life are excellent" results in a difficult-to-understand item of questionable validity.

Gadreau and Braaten (2016) and Gaudreau et al. (2015) also adapted an existing questionnaire to sport. They measured sport satisfaction using the eight-item subscale of the Multidimensional Student's Life Satisfaction Scale (Huebner et al., 1998), with the items adapted for sport (e.g., "My sport is interesting" instead of "School is interesting"). As can be seen from this example, Gaudreau's approach captures aspects that are rather circumstantial to the construct of satisfaction, such as interest, enjoyment (i.e., "I enjoy school activities"), positive/negative affect (e.g., "I feel bad at school"), and perceived competence/outcome (i.e., "I learn a lot at school").

Against this background, it becomes apparent that there is a need for a brief instrument to measure satisfaction with the domain of sport among individuals participating in various sports. One such German-language instrument is proposed in this study. It is an adaptation of an existing instrument capturing domain-specific satisfaction: the domain-specific self-satisfaction scale (Bereichsspezifische Selbstzufriedenheit, BSZ; Hormuth & Lalli, 1988, 1997). The BSZ is included in the Compilation of Social Science Items and Scales (Zusammenstellung sozialwissenschaftlicher Items und Skalen, ZIS), an open-access repository for social and behavioural science measurement instruments in German and English, published by GESIS – Leibniz Institute for the Social Sciences (https://zis.gesis.org/).

Table 1

Sport Satisfaction Scale (SSS; Sportzufriedenheitsskala, SZS) Items: Wording, Descriptive Statistics, and Inter-Item Correlations

Item	Original version of the German-language items (ad-hoc translation in parentheses)	Category frequency (%)				*M*	*SD*	Correlations (*r*)				
		1	2	3	4			1	2	3	4	5
1	Ich kann mich in dieser Sportart verwirklichen. (I can fulfil my potential in this sport.)	0.7	6.7	37.5	55.1	3.47	0.65	—				
2	Diese Sportart befriedigt mich. (This sport is satisfying to me.)	0.4	1.4	27.4	70.9	3.69	0.51	.49	—			
3[a]	Ich hätte gerne eine Sportart, die mich mehr ausfüllt. (I would like a sport that fulfils me more.)	5.6	10.2	40.0	44.2	3.23	0.85	.45	.54	—		
4[a]	Diese Sportart entspricht mir nicht. (This sport does not suit me.)	1.1	2.1	17.5	79.3	3.75	0.54	.37	.40	.42	—	
5[a]	Ein Wechsel dieser Sportart würde mir guttun. (Changing sport would be good for me.)	1.4	6.0	35.1	57.5	3.49	0.67	.37	.45	.51	.43	—

All items are rated on a 4-point scale (1 = *does not apply at all*, 2 = *does rather not apply*, 3 = *does rather apply*, 4 = *does fully apply)*. Suggested instruction: *To what extent do the following statements apply to you and the sport you practice?* Suggested original instruction: *Wie sehr treffen die folgenden Aussagen auf Sie und die Sportart zu, die Sie ausüben?* Original response format: (1 = *trifft gar nicht zu*, 2 = *trifft eher nicht zu*, 3 = *trifft eher zu*, 4 = *trifft genau zu*. $N = 285$. $r_{crit(5\%;\ one\text{-}sided)} = .10$.

[a] Reverse-coded item.

Theoretical Background of the Original Scale and Its Adaptation

Grounded in social psychology (Rosenberg, 1981) and structural symbolic interactionism (Stryker & Statham, 1985), the BSZ includes scales to assess satisfaction with five life domains: work, family and partner relations, social relations, and property. The instrument performed well in psychometric evaluations, demonstrating reliability as well as factorial, construct, and criterion validity (Hormuth & Lalli, 1988, 1997). Given the similarity between work and sport, we adapted the five items of the scale Arbeit (Work) by replacing "work" with "sport" in the original items, thereby creating the Sportzufriedenheitsskala (SZS), which, for the benefit of non-German speakers, will henceforth also be referred to by its translated name: the Sport Satisfaction Scale (SSS). It could also be said with equal justification that we used the Study scale (Studium), because Hormuth and Lalli reformulated the items of the Work scale in a completely analogous way so that they were suitable for capturing an individual's satisfaction with his or her study situation (form BSZ-S for student populations and form BSZ-N for general adult populations). The items (rendered in Table 1 with an ad hoc translation into English) are to be answered with respect to the primary sport an individual practises.

The Present Study

The aim of the present study was to evaluate the psychometric quality of the German-language Sport Satisfaction Scale. To achieve this, we examined its (1) factor structure and (2) internal consistency; (3) tested measurement invariance with respect to several demographic and sport-related variables; and (4) investigated the associations between the SSS and related psychological constructs and external criteria.

The first and last research questions were approached in light of theoretical expectations. Regarding the dimensionality of the SSS items, we assumed unidimensionality, based on Hormuth and Lalli's (1988) findings for the original Work scale. The analyses of construct and criterion validity, however, were guided by rather general hypotheses because there are hardly any theoretical or empirical studies on sport satisfaction in the above-described sense. Specifically, we drew on the dualistic model of passion (Vallerand, 2015). It differentiates two forms of passion: harmonious passion, which is characterised by autonomous motivation that promotes positive outcomes such as satisfaction and well-being, and obsessive passion, which is characterised by compulsion and internal conflict. Given these two qualities, it was expected that harmonious passion would be strongly (positively) associated with sport satisfaction, while obsessive passion would not be correlated to the same extent or at all. Furthermore, in line with self-concept theory (Shavelson et al., 1976), we hypothesised positive associations between sport satisfaction and self-concept because the theory emphasises the role of domain-specific experiences and feedback in shaping self-perceptions. This is supported by studies from other domains of life, which show moderate correlations between job satisfaction and self-concept (e.g., Cowin et al., 2008). Finally, we expected a positive association between sport satisfaction and athletic identity. Earlier research that examined satisfaction with athletes' performance or success in sport found positive relationships with athletic identity (Brewer et al., 1999; J. Schmid & Seiler, 2003; M. J. Schmid et al., 2024). These studies assessed satisfaction using either a single-item measure (Brewer et al., 1999) or adapted other existing measures to sport (e.g., The Life Satisfaction Scale of Ferring et al., 1996; J. Schmid & Seiler, 2003). Brewer et al. (1999) reasoned that athletes might dissociate themselves from their athletic role to make seasonal goals less relevant to self-evaluation and, consequently, to protect themselves from the disappointment that can accompany unsatisfactory sporting outcomes. Burns et al. (2012) argued, however, that differentiated findings are to be expected when the data are analysed at the facet level, that is, on the basis of the various dimensions of athletic identity. Because Burns et al.'s (2012) empirical study relied on the original – unidimensional – version of the *Athletic Identity Measurement Scale* (AIMS; Brewer

et al., 2022), it was not possible to derive any specific hypotheses for our study, which uses the current expanded version (AIMS-3G; Brewer et al., 2022) containing four additional dimensions.

The hypotheses we pursued as part of the criterion validity assessment were also informed by the basic psychological needs theory (e.g., Deci & Ryan, 2000) and empirical findings on the positive and negative effects of need satisfaction and frustration (for a summary, see Standage, 2023). The general hypothesis was that satisfaction with a sport is related to an individual's commitment to that sport. More specifically, satisfaction was expected to positively associate with five facets of (primary) sport commitment: weekly training volume, duration of involvement, participation in organised sport, and participation in competitions.

Methods

Participants

In spring 2024, a convenience sample of 365 German-speaking sports science students enrolled in three different bachelor's and master's courses at the University of Bern, Switzerland, was sent a link to an online survey. A total of 69 students either abandoned the survey before reaching the section on sports satisfaction or chose not to participate at all. With a response rate of 78.1%, the final sample comprised 285 students (M_{age} = 23.1 years; *SD* = 3.5; 43.9% female, 55.8% male, 0.4% other). Overall, 98.2% of the participants pursued extracurricular sports across 58 disciplines. The most popular activities were football, fitness, volleyball, gymnastics, jogging, athletics, and floorball. On average, participants had engaged in their primary sport for 10.0 years (*SD* = 5.9) and trained 4.1 hours per week (*SD* = 4.1). Likewise, 63.5% were organised in sports clubs, and 50.9% participated in competitions. (i.e., 53.4% at regional, 31.8% at national, and 15.7% at international level).

The authors' institutional research ethics committee approved the study. All respondents provided written informed consent to participate.

Measures

In addition to the different scales described below, standard demographic and sport-related information was collected: participants' gender, age; primary sport, number of weekly training hours and years of involvement (practice) in primary sport, participation in organized sports (basically club membership, 0 = *no*, 1 = *yes*), participation in competitions (0 = *no*, 1 = *yes*), and – if applicable – level of competition (1 = *regional*, 2 = *national*, 3 = *international; 0 = no competitive sport*).

Sport Satisfaction Scale

Sport satisfaction was assessed using the newly developed Sport Satisfaction Scale, adapted from the Domain-Specific Self-Satisfaction Scale (Bereichsspezifische Selbstzufriedenheit; Hormuth & Lalli, 1988, 1997). For a more detailed description of the theoretical background of the original scale, see the corresponding section in the Introduction. The scale comprises five items that are rated on a 4-point scale (1 = *does not apply at all*, 2 = *does rather not apply*, 3 = *does rather apply*, 4 = *does fully apply*). Example items are: "This sport is satisfying to me" ("Diese Sportart befriedigt mich"), "I would like a sport that fulfils me more" ("Ich hätte gerne eine Sportart, die mich mehr ausfüllt."), and "This sport does not suit me" ("Diese Sportart entspricht mir nicht"). Study participants were asked to indicate their satisfaction with their primary sport (for the remaining items, see Table 1). For the reliability of the scale, please refer to the Results section.

Passion Scale

Passion was assessed with the Passion Scale (PS; Vallerand et al., 2003) in its German version (Moeller, 2014). Harmonious Passion (7-items, e.g., "This activity is in harmony with the other activities in my life"; McDonald's ω = .75, 95% CI [.71, .80]) and Obsessive Passion (7-items, e.g., "I have almost an obsessive feeling for this activity"; ω = .88, 95% CI [.86, .90]) were rated on a 7-point scale, ranging from 1 (*strongly disagree*) to 7 (*strongly agree*) scale. Respondents

answered with reference to their primary sport. (All reliability estimates provided in this section and subsequent sections are based on the extant sample.)

Physical Self-Concept Scales

Physical self-perceptions were assessed using the Physical Appearance and Sports Competence scales of the Physical Self-Concept Scales (PSK; Stiller et al., 2004), developed for German-speaking populations and based on Marsh et al.'s (1994) Physical Self-Description Questionnaire (PSDQ). Participants responded on a 4-point scale (1 = *disagree* to 4 = *agree*). The Physical Appearance scale (9 items, e.g., "I am proud of my body") demonstrated high reliability (ω = .89, 95% CI [.87, .91]), as did the Sports Competence scale (6 items, e.g., "I am better at sports than most of my friends"; ω = .84, 95% CI [.79, .87]).

Athletic Identity Measurement Scales

Athletic identity was measured with the third-generation Athletic Identity Measurement Scales (AIMS-3G; Brewer et al., 2022; German-language version: M. J. Schmid et al., 2025). The AIMS-3G consists of (a) an Athletic Identity scale (4-items, e.g., "I would describe myself as an athlete"; ω = .91, 95% CI [.88, .92]), (b) two identity properties scales (i.e., Prominence, 5 items, e.g., "Sport is the most important part of my life"; ω = .92, 95% CI [.90, .93] and Self-Worth Contingency, 4 items, e.g., "My fitness level has a strong influence on how I feel about myself"; ω = .83, 95% CI [.79, .86]), and (c) two identity processes scales (i.e., Self-Presentation, 4 items, e.g., "It is important that I look like an athlete to others"; ω = .88, 95% CI [.85, .90]), and Social Reinforcement, 4 items, e.g., "Most of my friends are athletes"; ω = .72, 95% CI [.65, .77]). All items were rated on a 7-point Likert-type scale from 1 (*strongly disagree*) to 7 (*strongly agree*).

Procedure

The online questionnaire, programmed and administered in German via LimeSurvey, was sent out in spring 2024. Participants were given six weeks to complete the survey and received up to two reminders during the data collection period.

Data Analysis

The dimensionality of the five SSS items was analysed via confirmatory factor analysis (CFA). The analyses of measurement invariance were based on both multigroup CFA and tests of differential item functioning (DIF), which can be conducted using multiple indicators, multiple causes (MIMIC) models (e.g., Morin, 2023). Construct validity was assessed using exploratory structural equation models (ESEM), more precisely: Set-ESEM, an extension to ESEM. In this framework, "two or more a priori sets of constructs are modelled within a single model such that cross-loadings are permissible within the same set of factors (as in Full-ESEM) but are constrained to be zero for factors in different sets (as in CFA)" (Marsh et al., 2020, p. 103). In the given context, the different sets reflect the latent sport satisfaction factor and nine latent factors of passion, physical self-concept, and athletic identity.

To assess the reliability of the SSS, McDonald's coefficient ω and Cronbach's coefficient α were utilised, whereas model-based latent and manifest scale correlations (*r*) were estimated as indicators of its construct validity.

To evaluate the criterion validity of the SSS, we referred to the results of the multigroup CFA and the DIF analyses mentioned above, depending on the scale level of the criteria used. With regard to the categorical variables of gender, participation in organised sports, and participation in sports competitions, latent mean comparisons were made. For the continuous criteria of age, weekly training volume, and duration of involvement in the primary sport, we inspected the standardised path coefficients.

All analyses relating to structural equation modelling (SEM) were performed in Mplus (version 8.2; Muthén & Muthén, 1998–2017) using the robust maximum likelihood estimator (MLR) to account for deviations of the data from normality. The Mplus input code for selected models is available in electronic supplementary material ESM1. Supplementary retrospective power analy-

ses (Kline, 2023) were conducted using R Statistical Software (R Development Core Team, 2025), specifically the lavaan (version 0.6.20; Rosseel, 2012), semTools (version 0.5.6; Jorgensen et al., 2022), MASS (version 3.61; Venables & Ripley, 2007), and pwr (version 1.3.0; Champely, 2020) packages.

Criteria for Assessing Statistical Parameters

CFA, multi-group CFA, MIMIC, and Set-ESEM models were assessed using traditional fit indices: the comparative fit index (CFI; values > .95 for excellent fit, > .90 for acceptable fit), the Tucker-Lewis Index (TLI; values > .95 for excellent, > .90 for acceptable fit), and the root mean square of error approximation (RMSEA; values < .06 for excellent fit, < .08 for acceptable fit; West et al., 2023). Measurement invariance was evaluated by comparing changes (Δ) in CFI, TLI, and RMSEA between models. Decreases in CFI and TLI of ⩽ .01 and increases in RMSEA of ⩽ .015 were taken to provide evidence that the more parsimonious model should be retained, and thus a sequence is invariant (Chen, 2007; Cheung & Rensvold, 2002). Changes in CFI were prioritised over changes in TLI to evaluate alternative models, because the CFI is more robust than TLI in invariance testing and is typically the primary index used in this context (along with RMSEA). In addition, when models are extremely well fitting (TLI > 1.0), the baseline is inflated, resulting in a degradation of fit, which is not reflected in changes of χ^2, CFI, or RMSEA.

In terms of factor and scale reliabilities, values of .70 and above for McDonald's ω and Cronbach's α were considered sufficient in the research context (Nunnally & Bernstein, 1994).

Data Processing

The data were screened for anomalies in processing time. Given the specific nature of data collection in the various courses, only exceptionally short processing times were considered potentially anomalous (*Mdn* = 14.6 minutes, interquartile range [IQR] = 6.8, range = 5.4–68.3). Boxplot analysis of questionnaire response times, stratified by sample group, revealed no outliers in the lower tail of the distribution, indicating consistent minimum response times across all participants. However, screening the data with respect to the 57 study variables revealed several outliers in terms of age (10 cases over 30 years old), weekly training volume (two cases over 20 hours), and duration of involvement in sport (18 cases over 20 years). Taking into account other variables in the dataset (notably athletic performance level and sport), these extreme values were deemed plausible. However, values above the 99th percentile for these variables were winsorised to reduce the potentially undue influence of outliers.

Complete data were available from 188 participants (66.0%), with the remaining 97 cases accounting for the 1.9% missing values. Regarding the five items relating to satisfaction with sport, the maximum percentage of missing data was 1.8%. Little's MCAR test was not feasible due to the large number of study variables. Missing values were handled under the missing-at-random (MAR) assumption using full-information maximum likelihood in Mplus.

Results

Item Descriptive Statistics and Inter-Item Correlations

Table 1 also presents descriptive statistics for the items constituting the SSS as well as the inter-item correlations. Consistent with patterns commonly observed in satisfaction items (particularly those related to voluntary activities) the distributions exhibited either positive or negative skew, depending on item polarity. The inter-item correlations were moderate to strong, with an average of r = .43 (after reverse-recoding negatively worded items).

Table 2

Measurement invariance of the 1-factor confirmatory factor analysis of the Sport Satisfaction Scale (SSS) analysed with respect to gender, participation in organised sport, participation in sport competitions, age, number of weekly training hours and years of involvement in primary sport

No.	Model	χ^2	*df*	*p*	CFI	TLI	RMSEA	RMSEA 90% CI		SRMR	CM	$\Delta\chi^2$	Δdf	*p*	ΔCFI	ΔTLI	ΔRMSEA
								LB	UL								
0.1	1-factor CFA (N = 285) [a]	2.275	4	.069	1.000	1.019	.000	0.000	0.069	0.014	0.2	2.202	1	.138	.000	–.013	.000
0.2	1-factor CFA	4.337	5	.502	1.000	1.006	.000	0.000	0.077	0.019	—	—	—	—	—	—	—
	Gender																
	Single-group analyses																
1.1	Women (n = 125)	0.794	4	.939	1.000	1.079	.000	.000	.030	.013	—	—	—	—	—	—	—
1.2	Men (n = 159)	4.998	4	.288	.993	.982	.040	.000	.132	.024	—	—	—	—	—	—	—
	Invariance (multi-group) models																
1.3	Configural	5.177	8	.739	1.000	1.030	.000	.000	.071	.020	—	—	—	—	—	—	—
1.4	Weak	9.616	12	.650	1.000	1.017	.000	.000	.071	.071	1.3	4.560	4	.334	.000	–.013	.000
1.5	Partial strong [b]	13.419	15	.570	1.000	1.009	.000	.000	.072	.081	1.4	4.139	3	.247	.000	–.012	.000
1.6	Strict	12.200	20	.909	1.000	1.033	.000	.000	.029	.130	1.5	0.834	5	.975	.000	.024	.000
1.7	Latent variances/covariances	11.891	21	.943	1.000	1.037	.000	.000	.013	.125	1.6	0.017	1	.898	.000	.004	.000
1.8	Latent means	12.064	22	.956	1.000	1.039	.000	.000	.000	.124	1.7	0.003	1	.953	.000	.002	.000
	Participation in Organised Sport																
	Single-group analyses																
2.1	Unorganised (n = 104)	4.461	4	.347	.992	.981	.033	.000	.155	.035	—	—	—	—	—	—	—
2.2	Organised (n = 181)	1.395	4	.845	1.000	1.036	.000	.000	.063	.014	—	—	—	—	—	—	—
	Invariance (multi-group) models																
2.3	Configural	5.553	8	.697	1.000	1.024	.000	.000	.076	.024	—	—	—	—	—	—	—
2.4	Weak	7.251	12	.841	1.000	1.031	.000	.000	.050	.058	2.3	1.758	4	.780	.000	.007	.000
2.5	Strong	14.199	16	.584	1.000	1.009	.000	.000	.069	.082	2.4	8.739	4	.068	.000	–.022	.000

No.	Model	χ^2	*df*	*p*	CFI	TLI	RMSEA	RMSEA 90% CI		SRMR	CM	$\Delta\chi^2$	Δ*df*	*p*	ΔCFI	ΔTLI	ΔRMSEA
2.6	Partial strict [c]	16.493	19	.624	1.000	1.010	.000	.000	.063	.108	2.5	2.337	3	.505	.000	.001	.000
2.7	Latent variances/covariances	17.944	20	.591	1.000	1.008	.000	.000	.064	.150	2.6	1.339	1	.247	.000	–.002	.000
2.8	Latent means	20.173	21	.510	1.000	1.003	.000	.000	.068	.135	2.7	2.569	1	.109	.000	–.005	.000
	Participation in Sport Competitions																
	Single-group analyses																
3.1	No competition (n = 142)	3.647	4	.456	1.000	1.006	.000	.000	.122	.023	—	—	—	—	—	—	—
3.2	Competition (n = 143)	2.423	4	.658	1.000	1.040	.000	.000	.100	.022	—	—	—	—	—	—	—
	Invariance (multi-group) models																
3.3	Configural	6.127	8	.633	1.000	1.020	.000	.000	.082	.023	—	—	—	—	—	—	—
3.4	Weak	10.522	12	.570	1.000	1.010	.000	.000	.077	.096	3.3	4.309	4	.366	.000	–.010	.000
3.5	Partial strong [d]	15.963	15	.385	.996	.995	.021	.000	.083	.113	3.4	6.276	3	.099	–.004	–.015	.021
3.6	Partial strict [e]	20.092	19	.389	.995	.995	.020	.000	.077	.265	3.5	4.173	4	.383	–.001	.000	–.001
3.7	Latent variances/covariances	21.290	20	.380	.995	.995	.021	.000	.076	.320	3.6	1.158	1	.282	.000	.000	.001
3.8	Latent means	25.203	21	.239	.982	.983	.037	.000	.084	.355	3.7	5.330	1	.021	–.013	–.012	.016
	DIF: Age, weekly training volume, and years of involvement in primary sport																
4.1	Null effects model	35.036	19	.014	.950	.934	.054	.024	.082	.080	—	—	—	—	—	—	—
4.2	Saturated model	2.509	4	.643	1.000	1.029	.000	.000	.072	.011	4.1	34.683	15	.003	.050	.095	–.054
4.3	Factors-only model	19.512	16	.243	.989	.983	.028	.000	.064	.032	4.2	18.162	12	.111	–.011	–.046	.028

No.	Model	χ^2	*df*	*p*	CFI	TLI	RMSEA	RMSEA 90% CI	SRMR	CM	$\Delta\chi^2$	Δ*df*	*p*	ΔCFI	ΔTLI	ΔRMSEA
4.4	Partial invariant duration of involvement, invariant age and weekly training volume [f]	14.231	15	.508	1.000	1.004	.000	.000	.053	.026	4.2	12.392	11	.335	.000	–.025	.000
5	Set-ESEM with covariates: Harmonious and Obsessive Passion; Sports Competence, Physical Appearance; Athletic Identity, Prominence, Self-Worth Contingency, Self-Presentation, Social Reinforcement [g]	2020.273	1291	< .001	.909	.895	.045	.041	.048	—	—	—	—	—	—	—	—

N = 285. CFI = comparative fit index. TLI = Tucker-Lewis index. RMSEA = root mean square error of approximation. SRMR = standardised root mean square residual. MC = Compared model. The robust maximum likelihood estimator (MLR) was used to account for deviations of the data from normality. Accordingly, scaled (corrected) model test statistics are reported (i.e., MLR-χ^2, robust CFI and robust RMSEA). Differences in the fit indices of CFI, TLI, and RMSEA were computed by subtracting the fit index value for the later model from the fit index value of the earlier model.

[a] Based on theoretical considerations (positive wording), the uniquenesses of item 1 and item 2 were a priori allowed to correlate in the 1-factor representation of the Sport Satisfaction Scale. [b] To reach partial strong invariance, the intercepts of item 3 were allowed to be variant in the two groups. [c] To reach partial strict invariance, the uniquenesses of item 3 and item 4 were allowed to be variant. [d] To reach partial strong invariance, the intercepts of item 3 were allowed to be variant. [e] To reach partial strict invariance, the uniquenesses of item 3 were allowed to be variant. [f] Based on inspection of the modification indices, the effect of the variable years of involvement in primary sport on item 5 was freely estimated to reach an acceptable (i.e., not meaningful) deterioration in model fit when moving from the saturated model 4.2 to the most invariant model 4.4. [g] Based on empirical results (modification indices) and theoretical considerations, five correlated uniquenesses were introduced (Sports Competence: item 4 and 5; Physical Attractiveness: item 6 and 9 as well as item 7 and 9 [parallel wording]; Passion Scale: Obsessive Passion item 1 and 2 [Alamer & Marsh, 2022], and Obsessive Passion item 3 and Harmonious Passion item 6 [parallel wording]).

Factorial Validity

Given that the SSS comprises items with affirmative and negative formulation and negatively and positively worded items, a method effect might be present in the scale's factor structure. Therefore, we allowed a priori the error terms of the two positively worded items (items 1 and 2) to correlate (Cole et al., 2007). The data supported the hypothesised 1-factor CFA solution, $\chi^2 = 2.275$, $df = 4$, $p = .689$; CFI = 1.000, TLI = 1.019, RMSEA = .000 (Model 0.1 in Table 2). When comparing this model with the model without correlated uniquenesses (Model 0.2), the chi-square difference test was not significant ($\Delta\chi^2 = 2.022$, $\Delta df = 1$, $p = .138$) and no substantial difference was found in terms of CFI and RMSEA. Accordingly, the standardised parameter estimate for the correlated uniquenesses was negligible ($r = .136$, $p = .086$). Table 3 presents the standardised parameter estimates of the retained (hypothesised) model and reveals reasonably high loadings ($.59 \leqslant \lambda \leqslant .75$), indicating that the factor was well defined. Furthermore, the result of Horn's parallel analysis suggested that the SSS was unidimensional.

Table 3

Standardised parameter estimates from the 1-factor confirmatory factor analytic representation of the Sport Satisfaction Scale (SSS)

Items	λ	δ
1	.591	.651
2	.699	.512
3	.754	.432
4	.589	.653
5	.669	.553
Composite Reliability		
ω	.80 [.75, .85]	
α	.79 [.73, .85]	

N = 285. The hypothesised 1-factor model (Model 0.1) was supported by the data, $\chi^2 = 2.275$, $df = 4$, $p = .685$; CFI = 1.000, RMSEA = .000, 95% CI [.000, .069], SRMR = .014. CFI = comparative fit index. RMSEA = root mean square error of approximation. SRMR = standardised root mean square residual. A review of the model residuals revealed an acceptable local fit of the model. ω = McDonald's ω. α = Cronbach's α. In square brackets the 95% confidence interval of the reliability estimate is reported. λ = standardised factor loading. δ = item uniqueness.

Factor and Scale Reliability

McDonald's ω was estimated at .80, 95% CI [.75, .85], Cronbach's α at .79, 95% CI [.73, .85]. An examination of the descriptive statistics for the manifest scale – representing the average of participant responses across the five items – reveals a clear right skew, consistent with the item-level findings ($M = 3.52$, $SD = 0.48$, skewness = −1.28, kurtosis = 2.44). However, this scoring approach is not recommended (Swami et al., 2023).

Measurement Invariance

Measurement invariance was assessed with respect to gender (Models 1.1–1.8 in Table 2), participation in organised sport (Models 2.1–2.8), and participation in sport competitions (Models 3.1–3.8). Goodness-of-fit statistics are provided in Table 2. When the 1-factor CFA model was estimated separately in women and

men (Models 1.2 and 1.3), results revealed in both samples an excellent fit to the data (CFI and TLI > .95; RMSEA < .06). Results supported the weak (Model 1.4), but not the strong invariance of the model (Model 1.5). Inspection of the modification indices associated with the failed strong invariance model suggested that constraints had to be relaxed in the intercept of one item across genders (item 3: "I would like a sport that fulfils me more"; women were less satisfied with their sport than men). The models of partial strong and strict invariance (Models 1.5 and 1.6) were supported by the data. Moreover, evidence was found for the invariance of both the latent variances/covariances and the latent means model (Models 1.7 and 1.8).

With respect to participation in organised sports – that is, whether the students participated in sport in the context of a club or independently – both the configural (Model 2.3) and the weak (Model 2.4) invariance models were satisfactory in terms of model fit and relative change in fit. Strict invariance, however, was not achieved ($\Delta\chi^2$ = 18.423, Δdf = 5, p = .002; ΔCFI = −.066, ΔRMSEA = .076), suggesting measurement bias. Partial invariance models were pursued by freeing the equivalence constraint on the uniquenesses of two items (item 3: "I would like a sport that fulfils me more" and item 4: "This sport does not suit me") in both groups. This relaxation led to negligible changes ($\Delta\chi^2$ = 2.337, Δdf = 3, p = .505; ΔCFI = −.000, ΔRMSEA = .000) when comparing the strong and partial strict models. The remaining models of latent variances/covariances (Model 2.7) and latent means (Model 2.8) invariance were supported by the data.

Table 2 also contains the findings on measurement invariance in terms of participation in competitive sport. CFA results (Models 3.1–3.8) supported the weak but not the strong invariance. Examination of the freely estimated intercepts from the model of weak invariance and of the modification indices from the failed model of strong invariance indicated that the lack of strong invariance was limited to item 3 ("I would like a sport that fulfils me more"). Freeing the intercepts across the two groups resulted in a satisfactory model of partial strict invariance. Subsequent analyses revealed no evidence for strict invariance. Inspection of the modification indices suggested that invariance constraints had to be relaxed in the uniquenesses of, again, item 3. The resulting model of partial strict invariance (Model 3.6) was supported by the data. The invariance of the latent variances/covariances was also supported but not that of the latent means (ΔCFI = –.013, ΔRMSEA = .016).

Differential Item Functioning

Measurement invariance assessment was complemented with analyses of potential monotonic DIF based on three continuous variables, namely age, number of weekly training hours, and years of involvement in the primary sport (Table 2). The MIMIC tests of measurement invariance revealed that the saturated and factors-only models (Models 4.2 and 4.3) exhibited a superior fit to the data in comparison to the null effects model (Model 4.1). These results indicate an association between at least some of the predictors and the responses to the five SSS items. The finding that the saturated model fit substantially better than the factors-only model (ΔCFI = –.011, ΔRMSEA = .028) suggests that the effects of the predictor variables on the individual items cannot be fully explained in terms of effects on the latent factor. In other words, there is non-invariance of intercepts and thus evidence of monotonic DIF. Evaluation of the modification indices made it clear that item 5 was most likely affected by DIF ("Changing sport would be good for me"). Hence, a partial invariant model (Model 4.4) was estimated in which this item was regressed on duration of sport involvement. As can be seen from the estimated (standardised) regression coefficient (std. est. = −0.124, SE = 0.055, t = −2.273, p = .023), study participants with a long-standing involvement in their primary sport tended to score lower on item 5, that is the longer they had been practicing their primary sport, the less likely they envisaged a change of sport. A comparison of this partial invariant model (Model 4.4) with the saturated model (Model 4.2) showed no degradation of model fit in terms of the CFI and RMSEA but in terms of the TLI (ΔCFI = .000, ΔTLI = –.025, ΔRMSEA = .000).

Furthermore, the chi-square difference test was non-significant, $\Delta\chi^2 = 12.392$, $\Delta df = 11$, $p = .335$. In sum, we take these findings to indicate that with respect to participation in competitive sport, there is some (mild) measurement bias in the Sport Satisfaction Scale.

Construct and Criterion Validity

To evaluate construct validity on the latent level, we used Set-ESEM to investigate associations between Sport Satisfaction and several validation measures included in this study, namely passion (Harmonious and Obsessive Passion), physical self-concept (Sports Competence, Physical Appearance), and athletic identity (five factors). As seen in Table 2, goodness-of-fit statistics for Model 5 were acceptable, although marginal in terms of TLI, $\chi^2(1291) = 2020.273$, CFI = .909, TLI = .895, RMSEA = .041. The latent correlations based on this model between the SSS and the validation instruments are provided in Table 4 (alongside the manifest correlations). As expected, Sports Satisfaction was positively associated with Harmonious and Obsessive Passion (latent correlations: $r = .68$ and $r = .35$, respectively), and to a clearly smaller degree with Sports Competence ($r = .13$) and Physical Appearance ($r = .15$). Middling associations were observed between Sport Satisfaction and Athletic Identity, Prominence, Self-Worth Contingency, and Social Reinforcement ($.17 \leqslant r \leqslant .30$, $ps < .05$) but not Self-Presentation ($r = .05$).

Table 4

Construct validity analyses of the Sport Satisfaction Scale (SSS): Model-based latent and manifest scale correlations (r)

Measure	Sport Satisfaction	
	Latent	**Manifest**
Passion (PS)		
Harmonious Passion	.677	.592
Obsessive Passion	.353	.301
Physical Self-Concept (PSK)		
Sports Competence	.125	.124
Physical Appearance	.153	.132
Athletic Identity (AIMS-3G)		
Athletic Identity	.277	.288
Prominence	.247	.202
Self-Worth Contingency	.166	.115
Self-Presentation	.051	.050
Social Reinforcement	.304	.193

$N = 285$. $r_{crit(5\%;\ one\text{-}sided)} = .10$.

The above measurement invariance analyses also yielded latent variance/covariance and latent mean estimates for the two gender groups and the two groups defined by participation in organised sports. In particular, analyses revealed that women tended to have a lower latent mean on the Sport Satisfaction factor than men; however, the difference (which represents a standardised mean difference and is thus equivalent to Cohen's *d*) was neither statistically nor practically relevant (0.009 *SD*, $p < .952$). With respect to participation in organised sports (Model 3.8), individuals who were involved in organised sports tended to score higher than those who were not, but not significantly so (0.229 *SD*, $p = .110$). However, compared to individuals who took part in sports competitions, those who did not reported significantly lower levels of sport satisfaction (−0.324 *SD*, $p < .019$), which corresponds to a weak effect (Cohen, 1988).

The DIF analysis (Model 4.4) showed that age (under statistical control of the other effects) had no significant effect on the Sport Satisfaction factor (stand. path coefficient γ = .03, t = 0.351, p = .726). However, weak to moderate effects were observed for weekly training volume (γ = .20, t = 3.426, p = .001) and duration of involvement in the primary sport, in the sense that, as expected, greater time involvement in sport (γ = .18, t = 2.713, p = .007) was associated with higher sport satisfaction.

Post Hoc Power Analysis

Given that SEM typically requires large sample sizes, the validity of our results – based on a sample of 285 participants – may be limited. To address this, we conducted retrospective power analyses.

First, we conducted a MacCallum RMSEA-based power analysis (MacCallum et al., 1996) using the semTools R package for the one-factor CFA model (Model 0.1, df = 4). Under the close-fit hypothesis ($RMSEA_0$ = .05) versus a not-close alternative ($RMSEA_a$ = .08), power (1 – β) was low (.23) for our study sample. The estimated sample size required to adequately detect a lack of close fit was substantially larger ($N \approx 1803$).

Second, for multigroup CFA models (Models 1.1–1.8), a Monte Carlo simulation (500 replications) was implemented in lavaan and MASS R packages. Using observed loadings, residuals, and sample sizes (125 women, 159 men), we estimated post-hoc power for χ^2-square tests of measurement invariance. Power was low for most steps (metric: .16, scalar: .16, strict: .15, variance–covariance: .04) but higher for latent means (.76). Similar low-power patterns emerged for the other grouping variables – participation in organised sport (Models 2.1–2.8) and participation in competitive sport (Models 3.1–3.8) – so detailed results are omitted.

Third, we conducted an analogous supplemental simulation-based post hoc power analysis using lavaan for the MIMIC DIF analysis (Models 4.1–4.4). Results showed high power to detect small age-related DIF (.85–.90), moderate power for duration of involvement in sport (.55–.58), and lower power for weekly training volume (.27–.32).

Finally, a post hoc power analysis of manifest correlations (α = .05, β = .80, one-sided) indicated that the required sample size for detecting small, medium, and large effects was 617, 67 and 23, respectively.

While these analyses provide context, it should be noted that our model evaluation primarily relied on CFI and RMSEA (or changes therein), rather than chi-square or chi-square difference significance. Thus, these post-hoc power estimates are not critical and do not alter the substantive interpretation of our findings.

Discussion

In the present study, we adapted the SBZ Work scale developed by Hormuth and Lalli (1988, 1997) to the sport domain and examined the psychometric properties of the resulting German-language Sport Satisfaction Scale (Sportzufriedenheitsskala).

Regarding factorial validity, our results supported the hypothesised 1-factor representation. In anticipation of a potential method factor, we allowed the uniquenesses of the two positively worded items to correlate – a practice which may be viewed critically, notably when based on post-hoc indices of model fit (Landis et al., 2008). However, a subsequent analysis of the model without error covariance demonstrated that the two models were essentially identical, both in terms of model fit and factor loadings. Thus, concerns about inflated estimates due to unjustified specification of correlated uniquenesses seem to be unwarranted. Unfortunately, a direct comparison between the results of Hormuth and Lalli (1988, 1997) and our own is not possible on account of the different data analysis methods that were used (principal component analysis [PCA] versus CFA). However, the loadings of the five items had a similar range, with Hormuth and Lalli's (1988) loadings being higher on average (by approximately .13) – which is to be expected given that PCA and CFA handle specific item variance differently. Regardless of the reliability measure used, McDonald's ω or Cronbach's α, the values for the SSS were some-

what lower than those for the original Work and Study scale from the SBZ-N and SBZ-S, respectively (α = .87 in two independent samples; Hormuth & Lalli, 1988, 1997). Nevertheless, in the research context, the reliability of the SSS is satisfactory (ω = .80, α = .79).

The assessment of measurement invariance is a critical component of an instrument's validation. When a scale behaves differently across respondent groups, measurement bias emerges, distorting results and their interpretation (Chen, 2008; Guenole & Brown, 2014). In the present study, we found support for weak, (partial) strong, (partial) strict, and latent variance-covariance invariances across gender, participation in organised sport, and participation in competitive sport. Likewise, there was no indication of monotonic DIF with respect to age and weekly training volume. However, the intercept of one item ("Changing sport would be good for me") was not invariant across different levels of duration of involvement in sport, with experienced athletes tending towards lower item scores. From the perspective of identity theory and self-determination theory in sports, this lack of invariance may reflect the desire to maintain athletic identity, perceived competence, or social relatedness. Continuing in the same sport could be motivated by the need to preserve these aspects, which might be jeopardized by changing sports—especially after prolonged involvement. Overall, these findings indicate that it is empirically valid to compare latent SSS means between women and men, as well as between subgroups of different ages (in young adulthood) and various sport involvement variables, at least within the context of the present study.

The presented (cross-sectional) findings on construct validity support our theoretical expectations and suggest that sport satisfaction may indeed act as both determinant and outcome of sporting activities.

First, the results are consistent with the dualistic model of passion (Vallerand, 2015) in that passion towards a sport activity was differentially associated with satisfaction as a function of the type of passion underlying the engagement in a sport: Harmonious passion for a sporting activity was closely related to sport satisfaction; obsessive passion, in contrast, characterised by compulsion and internal conflict, was still positively associated with sport satisfaction but clearly less so. Thus, our study corroborated Vallerand et al.'s (2008) findings, also in the sport context, with respect to harmonious passion. With respect to obsessive passion, however, they did not find a path that predicted satisfaction. We see an explanation for this divergence in the fact that Vallerand et al. (2008) assessed life satisfaction, a construct that is broader than sport satisfaction. Moreover, obsessive passion has been shown to be conceptually closer to exercise addiction and pressure than to positive affective outcomes such as satisfaction (e.g., Paradis et al., 2013). Accordingly, the small effect size in our results appears theoretically plausible, because obsessive passion can increase commitment to sporting activities, but does so in a rigid and conflict-laden way that does not necessarily contribute to satisfaction.

Second, in line with Sonstroem and Morgan's (1989) exercise and self-esteem model, sport satisfaction fosters sporting activity and, in turn, subjective (and objective) sports competence. By the same token, our results support Wilson and Rodgers' (2007) proposition based on self-determination theory that satisfaction in sport settings contributes to positive self-perceptions. However, if it is assumed that the SSS at least partially captures the satisfaction of the three basic psychological needs (competence, autonomy, and relatedness), a stronger association with sport competence should be observed.

Third, it was not possible to formulate detailed hypotheses about the link between sport satisfaction and athletic identity. Nevertheless, our analyses revealed a substantial correlation between Sport Satisfaction and Athletic Identity. Sport Satisfaction was also associated with Prominence, Self-Worth Contingency, and Social Reinforcement – facets which were recently incorporated in the most prominent measure of athletic identity (AIMS-3G; Brewer et al., 2022) but have hardly been researched yet. At least post hoc, it seems plausible that if a person's athletic identity is prominent (with respect to other aspects of self-identity) and a person's self-worth is contingent on sport

performance and success, these athletic identity properties are likely to be nurtured by satisfactory experiences in their sporting endeavours (Brewer et al., 2018).

Regarding criterion validity, our hypotheses were largely supported: individuals who had trained extensively, been involved for a long time, and participated in competitions reported higher satisfaction with their primary sport. Conversely, those who were dissatisfied with their sport and might have considered taking up another one tended to be less committed and avoid participating in competitions. Contrary to expectation, participation in organised sport was not associated with satisfaction. It could be argued that belonging to a sports club is not a prerequisite for participating in sport, and that in a society that values individuality, it is not necessarily an indicator of sporting engagement and therefore may not be associated with satisfaction. However, it should be emphasised that we employed a cross-sectional design, which does not permit conclusions regarding causal direction. Future research should therefore utilise longitudinal designs and appropriate analyses (e.g., cross-lagged analyses) to examine these questions more rigorously.

Limitations and Conclusion

The present study provides evidence that the adaptation of the scale Arbeit (Work) from Hormuth and Lalli's (1988, 1997) German-language questionnaire on domain-specific self-satisfaction has been successful with regard to the primary criteria for psychometric quality, notably reliability, factorial validity, criterion validity, and construct validity. Special mention should be made of the extensive analyses of the instrument's measurement invariance. The results indicate that the SSS may readily be used across groups of gender, participation in organised sport, and participation in sport competitions. In addition, no indication of measurement bias was identified in relation to age or weekly training volume. A mild degree of invariance was, however, found for the duration of participation in sports.

Although these findings support the application of the SSS in a variety of contexts, as intended, several limitations of the study should be noted. First, due to practical constraints, we used a convenience sample of sports science students who were heterogeneous in terms of gender, sports practised and level of sporting involvement, but homogeneous in terms of education and age.

Second, because SEM is a technique designed for large samples, our sample of 285 sports science students was often not large enough for chi-square tests to detect small or moderate differences with adequate statistical power (e.g., .80). However, this is less relevant when model evaluation relies on fit indices such as CFI, RMSEA, ΔCFI, and ΔRMSEA, as was the case here. Furthermore, the observation of the hypothesised effects in the validation analysis suggests that the study was sufficiently sensitive to detect effects of this magnitude. Nevertheless, we encourage researchers to cross-validate our findings using larger and more diverse samples.

Third, the MIMIC approach used here does not cover the full taxonomy of measurement invariance tests, as it focuses on strong invariance, assuming – but not directly testing – invariant factor loadings (Swami et al., 2023). This limitation may raise questions about the validity of manifest-level group comparisons using the SSS (Kline, 2023; for an opposing view, see Robitzsch & Lüdtke, 2023). Future studies could address this by applying item response theory (e.g., Embretson & Reise, 2013) to rigorously assess both uniform and non-uniform DIF. Such analyses would not only strengthen comparisons within the German-language SSS but also support cross-language calibration and international comparisons.

In conclusion, the Sport Satisfaction Scale (Sportzufriedenheitsskala), measuring satisfaction with an individual's practised sport, has psychometric properties that justify its use as an economic research instrument in sport contexts (such as leisure sport). Nevertheless, future research should aim to replicate the presented findings in independent samples of more diverse populations. If the proposed tool proves

useful, we invite the research community to make it available in other languages. In particular, our ad hoc English translation of the questionnaire should be validated in order to increase its range.

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Data availability statement

Data supporting the findings of this study are available from the corresponding author (MJS) upon request. The Mplus input code for the selected models is available as electronic supplementary material (ESM1) on OSF: https://osf.io/bzsju.